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Migration and Local Public Services*

By

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April 2001

Abstract

Using unique Swedish micro data we examine the impact of local public services on community choice. The choice of community is modeled as a choice between a discrete set of alternatives. The US literature has produced conflicting evidence with respect to the importance of local public services. We find a robust positive (negative) relationship between local public services (local income tax rates) and the residential choices of short-distance migrants (defined as those moving within a local labor market). However, local public characteristics are less important for migrants who entered from other local labor markets. Using information on subsequent mobility, we also investigate whether the last result is due to lack of information about the characteristics of the local public sector. The evidence suggests that this is not the case.

Keywords: Migration, local public services, Tiebout, discrete choice, repeat migration
JEL classification: D12, H31, H73, R23

* We thank Per-Anders Edin, Eva Johansson and seminar participants at Uppsala University and IFAU for useful comments. Fredriksson acknowledges financial support from The Swedish Council for Working Life and Social Research and Dahlberg from The Swedish Council for Research in the Humanities and the Social Sciences (HSFR).
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1. Introduction

Since there is no market for local public services, it is not obvious how to estimate preferences for these services. In the literature, there exist several approaches to this problem. These include the median voter model (e.g., Bergstrom and Goodman, 1973), survey data approaches (e.g., Bergstrom *et al*., 1982), hedonic price models (e.g., Rosen and Fullerton, 1977), and discrete choice approaches. There are very few applications of the discrete choice approach: Friedman (1981), Quigley (1985), Boije and Dahlberg (1997), and Nechyba and Strauss (1998) are the only studies that we know of.

The two most well known of these studies, i.e., Quigley (1985) and Nechyba and Strauss (1998), arrive at conflicting conclusions with respect to the impact of local public services on household community choice. According to Quigley the impact is negative, while Nechyba and Strauss find that it is positive.[1] However, they also applied widely different approaches. Quigley studied a sample of Pittsburgh *renter* households who had *moved* within the last year; Nechyba and Strauss examined the *stock of homeowners* residing in six New Jersey school districts. The difference in the underlying populations is likely to drive the divergence in the results – at least to some extent.[2]

The purpose of this paper is to reexamine the question of the importance of local public services for community choice. We do this by using high quality Swedish data for movers to the local labor market of Stockholm.

It is an open question whether one should use the stock of residents or recent movers to estimate the demand for local public services. Sizable adjustment costs suggests that stayers may be off their demand curve. Migrants, on the other hand, may have preferences that are not necessarily representative for the population. Greenwood *et al* (1991) show that the equilibrium assumption inherent in using the stock of residents may result in the underestimation of the value of local amenities such as public services. So, in this paper

[1] Friedman (1981) concludes that local public services are not very important for residential choice; also, there is an "anomaly" in the sense that property tax rates enter positively (and significantly) into the equation. Boije and Dahlberg (1997) find that local public services exert a positive influence on community choice.

[2] There are other differences between the two studies. Quigley modeled the choice of town, neighborhood, and housing as a sequential choice where the town was chosen in the first stage of the decision process. Nechyba and Strauss concentrate on the choice of school district. The underlying population in Friedman (1981) was the stock of residents in parts of the San Francisco Bay area. The underlying population in Boije and Dahlberg (1997) was those who had purchased a house in the local labor market of Stockholm; their methodology was similar to Quigley (1985).

we focus on movers since if there is a significant relationship between local public services and community choice we are most likely to find it in this category.

We do, however, consider different categories of movers. In particular, we differentiate between individuals who have moved long-distance, defined as a move across local labor markets, and those who have moved short-distance, defined as a move within the local labor market of Stockholm. Long-distance movers presumably move for very different reasons and may lack the information necessary to optimize with respect to local amenities. We also provide some evidence on whether the lack of information is an important issue by looking at repeat migration in our sample of migrants.

Swedish data are very suitable for the purposes of this paper. First, the quality of the data is exceptional. Second, local governments comprise a sizable fraction of aggregate economic activity in Sweden: in 1992, local government expenditure amounted to around 27 percent of GDP; by comparison, expenditures at the federal and local level in the US amounted to 15 percent (OECD, 1994).[3] Third, local governments have important responsibilities such as the provision of day care, education, elderly care, and social welfare services. Finally, local governments have a large degree of autonomy regarding spending, taxing, and borrowing decisions.

We have access to a unique individual data set – LINDA; see Edin and Fredriksson (2000). LINDA contains the characteristics of a large panel of individuals and is representative for the Swedish population. From these data we have selected all individuals who moved to a new municipality within the local labor market of Stockholm between 1990 and 1991. To these data we match a set of (destination) characteristics of the local public sector and other characteristics of the municipality, such as housing.

Our results can briefly be summarized as follows. We find a robust positive (negative) relationship between local public services (local income tax rates) and the residential choices of short-distance migrants. However, local public characteristics are less important for migrants who entered from other local labor markets. Although there is more secondary mobility among long-distance movers, they do not move more within Stockholm relative to short-distance migrants. Thus, lack of information about the characteristics of

[3] In 1992, the share of local expenditures in total public expenditures was approximately the same (around 42 %) in Sweden and the US.

the local public sector does not appear to be driving the lower impact of local fiscal variables for this group.

The remainder of the paper is organized as follows. In the next section we present the econometric framework. In Section 3 we describe the data more thoroughly. Section 4 presents the results and Section 5 offers concluding remarks.

2. Econometric framework

In this section we begin by presenting the problem facing an individual deciding in which community to reside; then we outline the econometric specification.

Consider an individual who is confronted with a discrete set of location alternatives (communities) within a local labor market. When maximizing over this discrete set of alternatives she takes the attributes of the communities into consideration. In the spirit of Tiebout (1956), we mainly have local public services (g_c) in mind when characterizing the attributes of the community (*c*). We assume that the choice of local labor market has been made in a prior stage. Also, we take housing tenure and size choices as given.[4]

The individual, *i*, has additively separable preferences over the consumption of public goods and private goods, x_{ic} (housing consumption is subsumed into x_{ic}). We assume that the utility function is given by

$$u_{ic} = a_c + z(x_{ic}) + m(g_c) + \varepsilon_{ic} \tag{1}$$

where a_c denotes community amenities distinct from local public services. The random component of (1), ε_{ic}, captures random preferences for the (*c*)th alternative. The individual budget constraint takes the form

$$y_i(1 - \tau_c) = \rho_c x_{ic} \tag{2}$$

where y_i denotes income, ρ_c the price of private goods, and τ_c the local income tax rate. Thus, local public services are financed by income taxes.[5]

[4] In the empirical section, we report some evidence on the latter assumption. It turns out that less restrictive assumptions regarding housing choices yield only minor changes of the results.
[5] In Sweden, 99.6 % of the taxes raised at the municipal level come from income taxation. Moreover, the local tax rate is proportional so there is not much abuse of reality in specifying the left-hand side of (2).

For estimation purposes, we assume that the functions $z(\cdot)$ and $m(\cdot)$ in (1) are logarithmic. So a stylized version of utility would be

$$u_{ic} = \beta_0 \ln y_i + \beta_1 \ln(1-\tau_c) + \beta_2 \ln \rho_c + \beta_3 \ln g_c + a_c + \varepsilon_{ic} \tag{3}$$

We assume that y_i is determined by choice of local labor market. Since we consider choice of community conditional on choice of local labor market, y_i does not vary by *c* and can hence be ignored. The utility actually observed is the maximum over the set of all possibilities and (in principle) the coefficients have the interpretation of marginal utilities.[6] When estimating the parameters of (3) we will sometimes decompose local public service into its component parts.

Given that the utility observed is the maximum over the set of alternatives and ε_{ic} is i.i.d. with the type I extreme value distribution, McFadden (1973) has shown that the probability that individual *i* chooses community *c* is given by

$$\Pr(i \text{ chooses } c) = \frac{\exp\left(a_c + \beta_1 \ln(1-\tau_c) + \beta_2 \ln \rho_c + \beta_3 \ln g_c\right)}{\sum_c \exp\left(a_c + \beta_1 \ln(1-\tau_c) + \beta_2 \ln \rho_c + \beta_3 \ln g_c\right)} \tag{4}$$

Equation (4) implies the independence of irrelevant alternatives (IIA) property. The IIA assumption is, in principle, testable using a test devised by Hausman and McFadden (1984).

3 Data

We use two categories of data in this study: (*i*) data on the characteristics of individual migrants; and (*ii*) data on the attributes of the communities. We describe these data in turn, beginning with migrants.

3.1 The characteristics of migrants

Individual data on migrants come from the data base LINDA; see Edin and Fredriksson (2000). LINDA is a large panel of individuals, which is representative for the Swedish

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[6] The simple model outlined here of course implies the restriction $\beta_1 = -\beta_2$. Given that we only have approximate measures of local prices, we choose to enter prices and taxes freely throughout.

population; it covers around 3 percent of the population. The information in LINDA primarily comes from two data sources: filed tax reports and population censuses.

From LINDA we extract those 20-65 year olds that moved to a different municipality between 1990 and 1991 and where the destination municipality was located in the Stockholm labor market. Altogether there were 2,018 such moves; 1,444 moved to another municipality within Stockholm (our definition of a short-distance move) and 574 entered from another local labor market (our definition of a long-distance move).

Table 1 presents descriptive statistics for three categories of individuals; the first column gives the means and (where appropriate) the standard deviations for short-distance movers, the second column presents descriptive statistics for long-distance movers, and, for comparative reasons, the last column gives the means and standard deviations for those individuals who did not move at all.

Table 1. Descriptive statistics: movers and stayers.

	Short-distance movers Mean (std.)	Long-distance movers mean (std.)	Stayers mean (std.)
Individual characteristics			
Female	.458	.498	.504
Age	31.6 (10.1)	30.1 (9.5)	40.9 (12.0)
Immigrant	.188	.206	.198
Post high school education	.294	.321	.283
Earnings (SEK 100)	1,418 (941)	1,000 (862)	1,501 (1,050)
Employed	.891	.760	.870
Unemployed	.026	.111	.020
Welfare recipient	.055	.145	.044
Subsequent mobility	.369	.466	.174
Household characteristics			
Size of household	1.44 (.90)	1.33 (.86)	1.99 (1.18)
Kids ≤ 15 years of age	.184	.167	.294
Household earnings (SEK 100)	1,760 (1,335)	1,200 (1,202)	2,335 (1,724)
House ownership	.253	.340	.369
Employed family members	.191	.108	.440
# individuals	1,444	574	27,121

Notes: Except for subsequent mobility, all characteristics refer to 1990. Employed = 1 if individual earnings were greater than one basic amount. Unemployed = 1 if the individual received UI or Cash Assistance during 1990. Welfare recipient = 1 if the individual received welfare during 1990. Subsequent mobility =1 if the individual moved again between 1991 and 1997. Households are defined for tax purposes, i.e., married individuals and cohabiting individuals who have children in common are defined as a household. Employed family members = 1 if there were employed family members in the household according to the above definition. Individuals who did not move house between 1990 and 1991 are defined as stayers.

Migrants in general tend to be younger than stayers. Moreover, they are members of smaller households. The previous labor market history is strikingly different for long-distance movers compared to short-distance movers and stayers. Long-distance movers earned 40-50 percent less than the other two categories; their employment rates were 11-13 percentage points lower; and welfare receipt was substantially more prevalent. This suggests, of course, that long-distance movers primarily entered Stockholm for labor market reasons. Previous work has shown that these two groups exhibit different behavior with respect to out-migration; see Westerlund (1995) and Widerstedt (1998) for work on Swedish data. In a similar vein, we note that long-distance movers are more likely to move again within six years after their original move. In sum, it is probably reasonable to estimate separate locational choice equations for long- and short-distance movers.

3.2 Municipal characteristics

Table 2 presents summary statistics for the municipalities in the sample. The data has been obtained from Statistics Sweden. To avoid simultaneity problems we use 1990 characteristics throughout. We use expenditure data to proxy for the quality of local public services. This is of course unfortunate, but data reflecting the quality of services is very seldom available. In fact, we know of no study where community choice has been related to the *quality* of public services.

Average total expenditure amounts to over 1,500 Million SEK, which corresponds to 165 Million PPP-adjusted US$ in 1990. Hence, by international standards the Swedish local public sector is large. The prime responsibilities of the municipalities are schooling and care for children and the elderly. Expenditures on child and elderly care include labor costs, rents, and administration costs. With respect to education expenditure, however, we are able to exclude rents and administration costs so that this item only includes expenditures related to teaching. Panel A of Table 2 shows that, on average, 13 percent of expenditure is devoted to teaching at the compulsory level and 32 percent is devoted to child and elderly care. The remainder of the local budget (55 percent on average) is allocated to culture, parks and recreation, high-school education, administration, and assistance programs such as social assistance (welfare) and housing assistance.[7]

[7] Ideally, we would have liked to separate expenditures on high-school education from those included in other expenditures. However, 5 out of the 22 municipalities for which we have disaggregate school expenditure data do not provide high school education; instead they buy these services from neighboring municipalities.

Panel B of Table 2 presents local variables as we introduce them in the empirical analysis (although we enter some variables in logs). Our general strategy is to measure each expenditure item per potential user; total expenditure is measured relative to the population, child care expenditure is measured relative to the size of the population aged 0-6 and so on. For estimation purposes, we note that there is a fair amount of variation in local expenditure. The coefficient of variation for the expenditure items ranges from 9 to 19 percent.

Table 2: Descriptive statistics, municipalities.

	Mean (std.)
A. Expenditure	
Total	1,541,007 (3,454,629)
Percent of total expenditure devoted to…	
…child care	24
…education (expenditures on teaching at compulsory level)	13
…elderly care	8
…other purposes	55
B. Variables relevant for the empirical analysis	
Total expenditure (per capita)	22.090 (2.690)
Child care (per individual aged 0-6)	56.188 (9.265)
Education (expenditures on teaching, per individual aged 7-15)	23.921 (2.266)
Elderly care (per individual aged 65--)	16.290 (3.136)
Other purposes (per capita)	12.226 (2.259)
Municipal tax rates (percent)	14.73 (1.24)
Social assistance (norm 1)	111.69 (7.50)
Social assistance (norm 2)	185.77 (9.66)
House price	1291.115 (447.741)
Vacant rentals	8.73 (23.61)
Population size	63,256 (125,843)
Share of foreign citizens (percent)	8.88 (3.84)
Municipal unemployment (percent of population age 18-65)	0.60 (0.20)
# Municipalities	26

Notes: Expenditures and house prices are expressed in thousands of SEK. The house price used is the average price of houses sold in a municipality in 1990. Social assistance (norm 1) is the municipality norm for single-person households (in percent of the basic amount). Social assistance (norm 2) is the municipality norm for married or cohabiting persons (in percent of the basic amount). Since we know the marital status of the households in our data, we can attach the appropriate social norm to each of the observations. This is what we have done for the variable “social assistance”, which is the variable we use in the estimations. Expenditures on teaching at compulsory level had to be imputed for four municipalities. The imputation procedure is described in Appendix A.2.

The bottom half of panel B reports some other characteristics that we will condition on in the empirical analysis. These characteristics include welfare generosity, some information pertaining to the housing market, population characteristics, and unemployment rates.

Municipalities are free to determine the generosity of social assistance (welfare); the Swedish system is thus similar to the American system in this respect. We report two measures of welfare generosity. The first measure (norm 1) pertains to singles, while the second (norm 2) pertains to married or cohabiting couples. A feature of our data is that we know the marital status of each person in our sample.[8] Hence we can assign the norm that is of relevance for the particular individual, yielding local and individual variation in welfare generosity. This is the approach we take during estimation and we normalize the norm by the number of adult members of the household.

The characteristics of the housing market are summarized by the average price on sold houses during 1990 and the number of vacant rental apartments in September of 1990. The three major tenure forms in the Swedish housing market are owner occupancy, condominiums (coop shares), and renting. These tenure forms accounted for 22, 17, and 47 percent, respectively, of the total number of apartments in the Stockholm area in 1990. The Swedish housing market is far from the idealized competitive one. This is particularly true for the rental market, where there are price restrictions and rationing rather than prices being determined by supply and demand. Thus, attractive areas feature longer queues rather than higher rents; in principle, there should be no price differences for dwellings of equal size and quality across the country. To capture the fact that the rental market exhibits rationing we use the number of vacant public rentals in the regressions. The number of vacancies was extremely low because of the booming housing market in 1990.

The bulk of regional price variation within the Stockholm area is due to house prices. Market forces essentially determine the prices of non-rental housing. However, there is only price information pertaining to owner-occupancy, which is directly relevant for only 22 percent of the market. Even if we make the assumption that the prices of "coops" are proportional to the prices of owner-occupied housing there is still 47 percent of the market where the price information is of limited relevance.

Given that we hold *all* regional amenities constant, we would like to think about higher house prices as a deterrent to entry. However, the assumption that we measure *all*

[8] Notice, though, that households are defined for tax purposes, meaning that cohabiting individuals who have children in common are classified as households (together with married individuals). Thus the number of cohabiting individuals is underreported in our sample.

regional amenities is not particularly realistic. Hence, the sign of house prices is ambiguous if there is some capitalization of amenities into prices (see e.g. Yinger, 1982, on the idea that local public services and taxes will be capitalized fully into house prices). Although the interpretation of the house price variable is problematic, capitalization has the virtue that there is less risk of misspecification in the sense that any relevant variable that we leave out of the model will to some extent be included if we control for house prices.

We consider two measures to control for population characteristics: population size and the share of foreign citizens. The municipalities of the Stockholm labor market vary substantially in size. The extreme case is the Stockholm municipality, which is 100 times larger than the smallest municipality (Vaxholm) and eight times greater than the second largest one. Thus, the largest share of the inflow will enter the Stockholm municipality by construction. To avoid these "mechanical" effects we control for population size throughout.

According to Table 1, around 20 percent of movers are foreign-born. In the literature on immigrants' internal migration, it has been shown that they are attracted to localities with large fractions of foreigners; see, e.g., Zavodny's (1998) survey of the US studies and Åslund (2000) on Swedish data. Therefore, it is potentially important to control for immigrant concentration. In general, of course, immigrant concentration may represent an attracting force for some and a repelling force for others.

The last regional characteristic that we consider is local unemployment. The unemployment to population ratio in Stockholm is extremely low in 1990. A long economic upturn starting in the beginning of the 1980s peaked around 1990; the aggregate unemployment to population ratio stood at 1.4 percent in 1990. If Stockholm truly is one single local labor market, then labor market prospects as such should not matter for community choice. However, it may still be the case that agents dislike (or like, depending on individual characteristics) living in unemployment-ridden communities.

4. Results

Having described the data, we now turn to the estimation results. In this section we ask questions such as: How important are local public services for choice of community? Does the importance vary across different categories of individuals? Why does the response to variations in local public services vary across categories?

We begin this section by analyzing the determinants of community choice (sections 4.1 and 4.2). Since we find that the response pattern differs between short- and long-distance movers, we proceed to analyze subsequent mobility (section 4.3). If we would find that mobility within the Stockholm labor market is greater among long-distance movers this would suggest that this category of movers were not able to optimize in the first stage – either because of lack of information or restrictions in the housing market.

4.1 Determinants of community choice: Baseline estimates

Our general strategy is to begin with a bare-bones specification that is consistent with equation (4). As we go along we add more sophistication in the sense of adding more controls. The baseline estimates are presented in Table 3.

Table 3: Logit results for choice of municipality.

Variables	**I** (All movers)	**II** (All movers)	**III** (Short-distance)	**IV** (Long-distance)
Expenditure (in logs)				
Total	2.180	1.910	2.447	.303
($\ln(g)$)	(.268)	(.342)	(.395)	(.713)
"Prices" (in logs)				
Tax retention rate	11.378	15.938	17.319	10.362
($\ln(1-\tau)$)	(2.808)	(3.108)	(3.464)	(7.196)
House price	–.049	–.177	–.269	.253
($\ln(\rho)$)	(.142)	(.155)	(.173)	(.360)
Other variables				
Population size ($\times 10^{-5}$)	0.349	.377	.347	.442
	(0.012)	(.014)	(.017)	(.028)
Vacant rentals ($\times 10^{-2}$)		.377	.426	.131
		(.115)	(.127)	(.285)
Social assistance		.005	–.007	.037
		(.006)	(.006)	(.012)
Share of foreign citizens		5.100	3.696	9.643
		(1.085)	(1.227)	(2.365)
Local unemployment		–48.373	–87.435	67.663
		(19.832)	(23.359)	(39.479)
Pseudo R-squared	0.193	0.196	0.146	0.356
Log L	–5,308	–5,286	–4,020	-1,203
# individuals	2,018	2,018	1,444	574

Notes: Standard errors in parentheses.

Column I of Table 3 reports the coefficient estimates from a baseline specification, including only (the log of) total expenditure, (the log of) the tax retention rate ($1 - \tau$), (the log of) house prices (as a proxy for ρ), and population size (to control for the mechanical effects of size). The results clearly show that individuals opt for municipalities that offer a high level of *per capita* expenditure given taxes; analogously, individuals move to municipalities offering lower tax rates given local public expenditure. In the stylized framework of section 2, the ratio of the two coefficients is related to the marginal rate of substitution between public and private goods (i.e. net income); according to the estimates of column I, agents require an income increase of around 0.19 percent to compensate for a reduction of public services of one percent. High house prices do not seem to deter individuals from entering a municipality. This result is broadly consistent with the idea that regional amenities capitalize into house prices;[9] however, it may also reflect the fact that we measure a price that is of limited relevance for a substantial fraction of the sample.

Column II adds a set of other characteristics that are potentially relevant for community choice. Local public services and tax rates are still highly significant, but the addition of controls shifts the relative importance of local tax rates and local public services somewhat. The compensating variation for a reduction of public services by one percent is 0.12 percent of net income according to the estimates in column II.

Most of the coefficients on the additional controls are in line with expectations: housing vacancies attract migrants, while poverty stricken areas, as measured by unemployment, deter migrants. Welfare generosity enters insignificantly, which is not too surprising given that generous welfare benefits may represent incentives for some migrants and disincentives to others. What is more surprising, perhaps, is the fact that immigrant concentration enters positively and significant. This is due to the fact that the foreign-born constitutes close to 20 percent of the sample, and they value immigrant concentration positively.[10]

[9] Suppose that local public services and taxes are the only regional amenities and that they are fully capitalized into house prices. Then all three coefficients are not simultaneously identified: either the coefficients on taxes and public expenditure or the coefficient on house prices would drop out of the equation.

[10] For short-distance movers, a simple interaction between immigrant status and the share of foreign citizens suggests that the positive coefficient is mostly driven by the foreign born; the coefficient for natives is positive but insignificant.

Columns III and IV split the sample into short-distance and long-distance movers. While the coefficients on public sector characteristics are well determined for short-distance movers, this is not the case for long-distance movers. The estimates on local tax rates and public services are smaller in size and insignificant in the latter category. The sole public sector characteristic that is a significant determinant of community choice for long-distance migrants is welfare generosity. This is in line with the observation that those who have entered from another labor market are from the lower end of the earnings distribution; see Table 1.[11] Also, the fact that higher unemployment attracts long-distance migrants is an odd feature of these estimates, although the effect is only statistically significant at the ten percent significance level.

The difference between columns III and IV with respect to public sector characteristics may have several explanations. One hypothesis is that the long-distance movers lacked the knowledge to make an informed choice with respect to, e.g., public services. Another hypothesis is that they had the information, but their primary concern was finding an affordable apartment in order to make the move to, e.g., a new job.[12] Both of these explanations suggest that long-distance movers would move again (within Stockholm) in order to correct a sub optimal initial choice. We examine this prediction in detail below.

The results in columns III and IV thus imply that in order to get reliable estimates of the preferences for local public services one should focus on short-distance migrants. Accordingly, we take this approach when examining the determinants of community choice further.

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[11] This result is broadly consistent with the US evidence presented in Meyer (1999). Meyer finds a small positive coefficient on welfare generosity for single mothers.

[12] This explanation could rationalize the coefficient on the local unemployment rate for long-distance movers. In high unemployment communities the housing market is relatively depressed making it easier to find an affordable apartment, in particular for individuals at the lower end of the earnings distribution.

4.2 Determinants of community choice: Further evidence

The purpose of this subsection is to present some further evidence on the determinants of community choice. The questions we examine are: Do preferences vary with respect to the different components of local public services? Do preferences vary by individual characteristics? We present the results of some variations of the baseline estimates in Table 4.

As a preliminary exercise, we test whether preferences differ by expenditure item. For that purpose we decompose total expenditure using the first-order approximation:

$$d\ln g_c \approx \sum_k \mu_c^k d\ln g_c^k \tag{5}$$

where μ_c^k is the share of expenditure item k in total expenditures. We attach a separate coefficient to each item and test the hypothesis that the coefficients are equal by a likelihood ratio test. Based on the results from this exercise (not shown), we decisively reject the null hypothesis.[13]

Since we reject the hypothesis of preference equality, we examine the issue further. We introduce each component in logs and measure expenditure relative to the number of potential users (see panel B in Table 2 for exact definitions);[14] column I gives the results. The two most important expenditure items – both in terms of size and significance – are education expenditure and expenditure on "other purposes". Spending on child and elderly care is comparatively less important.

In columns II and III, we investigate whether preferences for local public services vary by income and age.[15] In column II, we interact aggregate expenditures and local tax rates with earnings; in column III, we introduce similar interactions with age.

The results in column II suggest that the marginal utility derived from local public services is lower among high-income groups; the interaction between income and

[13] The test statistic for equality of the coefficients was 10.84 (three degrees of freedom) with an associated p-value of 0.013.

[14] This is not an exact decomposition of aggregate expenditures and hence cannot be used for purposes of testing. We choose to present the result from this specification as the coefficients are easier to interpret; the coefficient directly gives the response to a one percent change in the particular expenditure item. The coefficients from the specification implied by the decomposition in equation (5) have the interpretation of the response to a one percent change in aggregate expenditure stemming from a particular expenditure item.

[15] We have also tried interactions with other background characteristics. For instance, there was no significant difference between individuals who had children under the age of 16 and those who had not.

aggregate expenditures is negative and significant. This is consistent with the observation that municipalities are responsible for sizable welfare programs such as social and housing assistance to which the high-earners do not qualify.[16] Analogously, high-income groups are less prone to enter municipalities where the local tax rate is high.

Table 4: Logit results for choice of municipality. Short-distance movers.

Variables	**I**	**II**	**III**
Expenditure (in logs)			
Total: ($\ln(g)$)		3.352	4.926
		(.508)	(.764)
Interaction: $\ln(g)$×(earnings) (×10^{-2})		–.062	
		(.022)	
Interaction: $\ln(g)$×(age)			–.076
			(.020)
Child care	.374		
	(.333)		
Education	1.033		
	(.465)		
Elderly care	.312		
	(.234)		
Other purposes	1.150		
	(.313)		
"Prices" (in logs)			
Tax retention rate: ($\ln(1-\tau)$)	17.855	12.213	13.695
	(3.469)	(4.546)	(7.068)
Interaction: $\ln(1-\tau)\times$ (earnings) (×10^{-2})		.342	
		(.191)	
Interaction: $\ln(1-\tau)\times$ (age)			.120
			(.186)
House price: ($\ln(\rho)$)	–.205	–.264	–.260
	(.233)	(.173)	(.172)
Other variables			
Population size (×10^{-5})	0.349	.347	.347
	(.024)	(.017)	(.017)
Vacant rentals (×10^{-2})	.377	.431	.434
	(.135)	(.127)	(.127)
Social assistance	–.007	–.007	–.009
	(.007)	(.006)	(.006)
Share of foreign citizens	3.554	3.610	3.479
	(1.441)	(1.228)	(1.229)
Local unemployment	–94.434	–86.702	–86.395
	(26.112)	(23.384)	(23.443)
Pseudo R-squared	0.146	0.147	0.147
Log L	–4,019	–4,014	–4,012
# individuals	1,444	1,444	1,444

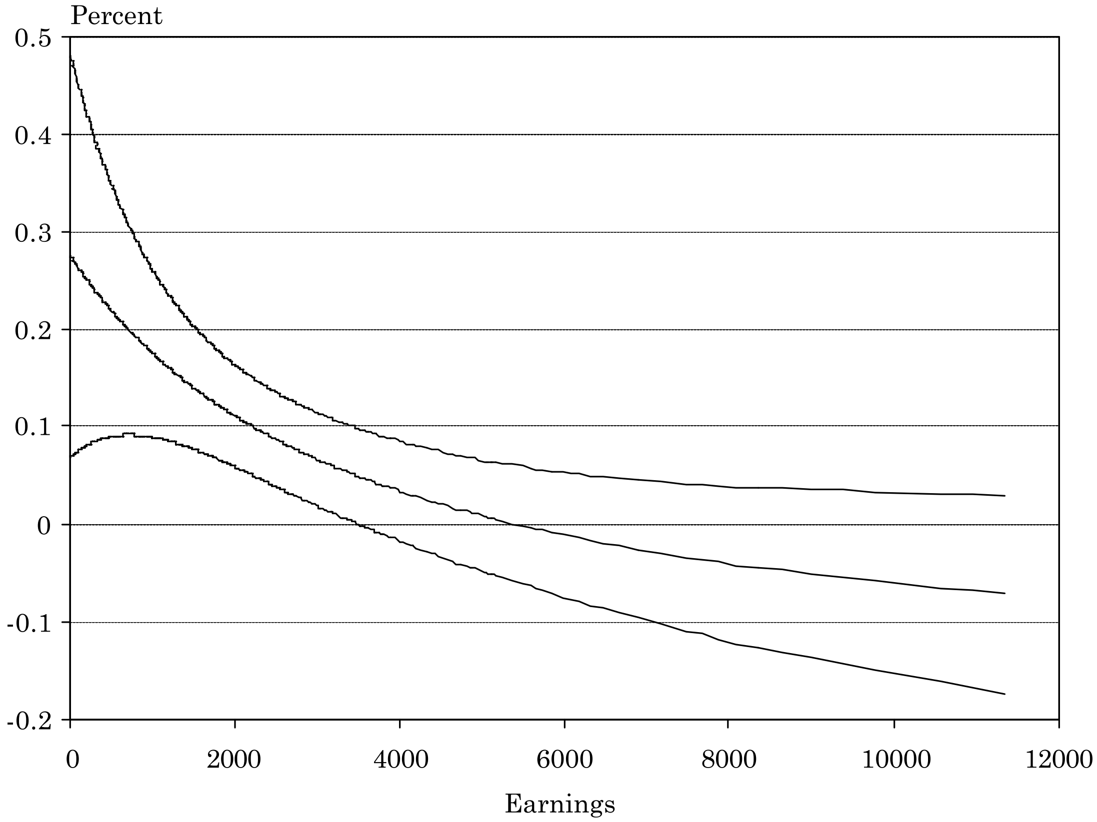
Notes: Standard errors in parentheses.

In Figure 1 we graph the compensating variation implied by these estimates. High-income individuals require a lower increase in net income to compensate for a reduction of public services. Notice, though, that there are some extreme cases that distort the picture

[16] Indeed, a specification with each expenditure item interacted with income suggests that "other purposes" (which includes all the welfare programs) is driving the negative interaction term for aggregate expenditure.

somewhat. The overwhelming majority of the sample places a positive value of the services provided by the local public sector; only 0.3 percent of the sample has a negative valuation of public services. Also, the negative compensating variations are not significantly different from zero.

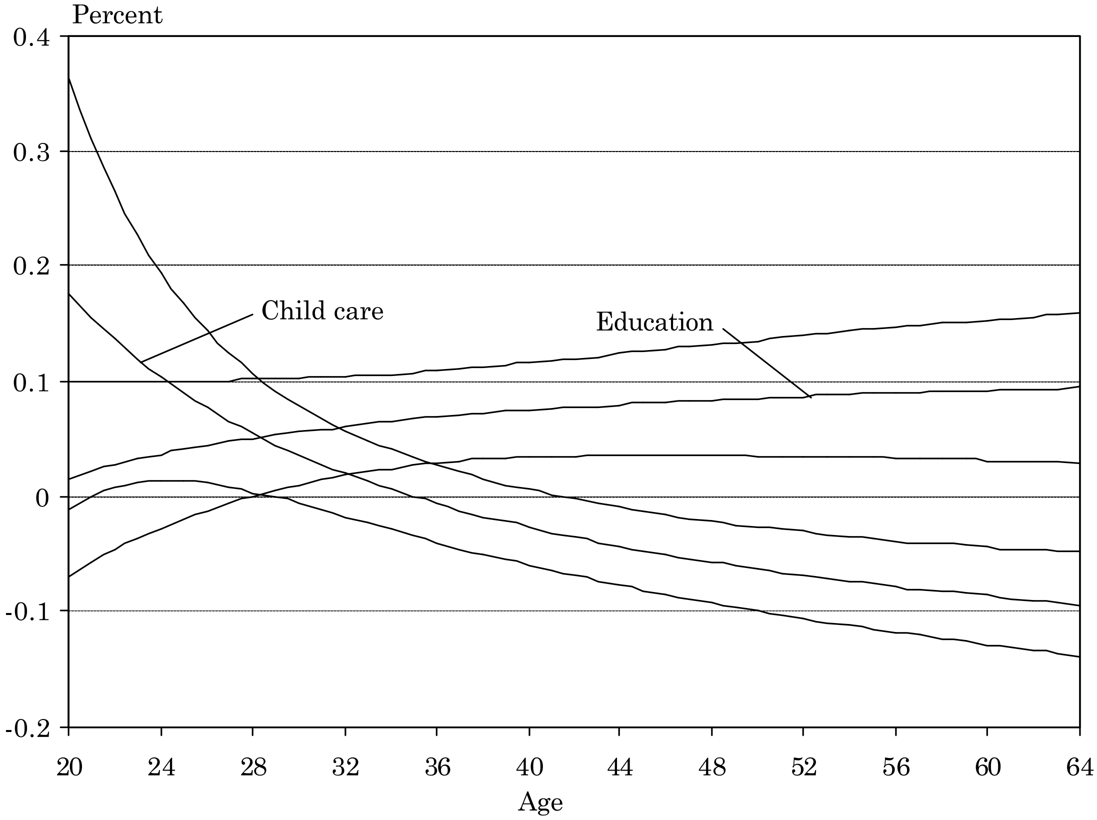
Figure 1: Compensating variation as a function of earnings.



Note: The graph plots the percent increase in net income required to compensate for a reduction of public services by one percent. The thin lines show the 95 percent confidence interval (calculated using the delta method).

The interactions with age in column III suggest a similar pattern, although the interaction with the local tax rate is insignificant. If we take the estimates at face value, we can transform the interaction estimates in column II and III such that the average valuation reflects the age and earnings distribution of the stayer population. The difference in the age distribution between the mover and stayer population matters most; the age distribution of the stayer population implies an average compensating variation of 0.10 percent, which should be compared to 0.15 percent for the mover population. However, the earnings distributions are so similar that there was not much of a difference between the two populations (0.15 % for stayers and 0.16 % for movers).

Figure 2: Compensating variations by age for education and childcare.



Note: The graph plots the percent increase in net income required to compensate for a reduction of each expenditure item by one percent. The thin lines show the 95 percent confidence interval (calculated using the delta method).

Rather than interacting age with aggregate expenditure, it is perhaps more interesting to examine whether the preferences for different kinds of expenditure varies with age. The basic results are presented in Figure 2; Table A1 reports the full estimation results.

The expenditure items with a significant age interaction are childcare and education expenditure; for this reason we only plot the compensating variations for these two items. Demand for childcare declines with age, while demand for education increases with age. These patterns broadly conform to intuition.[17] Comparatively young individuals are more likely to have children in kindergarten age and so value the increase in childcare

[17] That the compensating variation for a reduction in childcare expenditure is negative for individuals above 42 years-of-age is a counterintuitive feature of these estimates; also one would expect the compensating variation for a reduction of education expenditure to decline with age. Notice in these respects that there is good reason to treat the standard errors at the higher end of the age interval with more than the usual skepticism as only 20 % of the sample is above age 40. Since we are calculating asymptotic standard errors (using the delta method) we are imposing no penalty to having few observations at a particular age. Presumably, the variability at the higher end of the interval would be substantially greater had we, e.g., calculated the standard errors by means of bootstrapping.

expenditures more. Older individuals are more likely to have kids aged 7-15, and, hence, have a stronger preference for increases in school expenditure. Notice also that we have experimented with more flexible age interactions (we used a quadratic in age) without changing the overall flavor of the results.[18]

It may also be of interest to compare our estimates for educational expenditures with the US estimates in Nechyba and Strauss (1998). The response elasticity with respect to education is around 1 in our case; the corresponding number for the US is around 6.[19] The estimated elasticity is hence six times smaller for Sweden than for the US.[20] Of course, there might be several explanations for this – one obvious candidate is the difference in the underlying populations. However, we would also like to emphasize two of the institutional differences between Sweden and the US. First, the years we study feature a minimum school quality standard that was implemented by the central government via specific grants (the municipalities were free to improve on this standard by increasing education expenditure). Effectively, the binding minimum standard eliminates the lower tail of the school quality distribution and, given decreasing returns to school quality, the marginal utility of an increase in school expenditure may be lower than in a situation with no minimum standard.[21] Second, the Swedish school system is decidedly more egalitarian than the American one; see, e.g., Lindahl (2000). More resources are directed towards the less able in Sweden. Therefore, a given variation in education expenditure may be less related to variations in the quality of publicly provided education for the average individual.

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[18] All our estimates are based on the IIA-assumption. In principle, this assumption is testable using the test developed by Hausman and McFadden (1984). In practice, however, it is far from clear how to implement the test procedure – especially when the number of alternatives is as large as 26. The test is based on a set of alternative models where one or more of the choice alternatives are left out. In our setting, there literally exist an infinite number of possible alternative model specifications. To get a flavor of how well specified our models are, we conducted 26 tests of the model with age interactions (see column I in Table A1). In these tests, one community at a time was left out. Notice that, on the basis of probability calculus, we should expect the model specification to be rejected in 1 or 2 cases at the 5 percent level. In this example we rejected the model in 15 out of 26 cases, which suggests that the model may be incorrectly specified. Notice also that all models in Nechyba and Strauss (1998) are mis-specified despite their contention of the contrary.

[19] The reported elasticity refers to a marginal increase in educational expenditures in the chosen community. The elasticity is given by $(1 - P)\lambda$, where P is the probability of choosing a given community and λ is the coefficient on school expenditures. For simplicity, we use $P = (1/C)$, where C denotes the number of communities, in these calculations. In our case, the estimate is based on column I in Table 4; the US number is based on the estimates in Table 3, Model 1, in Nechyba and Strauss.

[20] These results are broadly consistent with the results in Ahlin and Johansson (2001); investigating the demand for local public school expenditures using survey data, they find that the income elasticity is higher in Sweden than in the US.

[21] Think of a world where parents have preferences defined over their own consumption and the consumption of their children. An increase in school expenditure raises the future consumption opportunities of children. The marginal utility of public education (for parents) will then be positively related to the return to an increase in public education.

Our analysis so far is based on the assumption that housing choices are given. What happens to the estimates if we relax this assumption? Suppose that we are willing to assume that utility is (weakly) separable in housing consumption (h), other private goods (x), and public services (g), such that we can solve the maximization problem in stages. Suppose further that we can treat housing as a one-dimensional continuous good. Consider the optimal choice of housing taking x and g as given. This problem would give housing demand function of the form: $h_c = h_c(y_i(1-\tau_c), p_c, x_{ic}, g_c)$, where p_c denotes the marginal price of an additional housing unit. Substituting the housing demand functions back into the original utility function we have what is sometimes referred to as a polytomous choice model; see Quigley (1976). The upshot of this model is that we should amend the original model with net income and a measure of the marginal price of an additional housing unit. Notice, though, that the coefficients from this "indirect utility" framework do not have the same interpretation as the original ones. For instance, the coefficient on local public services no longer reflects the marginal utility of public services but rather the combined effect of a change in public services and the induced change in housing consumption. Nevertheless, the model is less restrictive and we use it to assess the robustness of our results.[22]

With respect to public services, the results from the polytomous choice model are remarkably similar to the original estimates (the estimates are reported in Table A1). In the polytomous choice setting, the coefficient on the log of aggregate expenditures is 2.678 (standard error: 0.784), which should be compared to our previous estimate of 2.447; c.f. col. III in Table 3. The inclusion of individual income (interacted with a set of community dummies) reduces the size and significance of the tax retention rate. However, this is precisely what one should expect given that local tax rates and income is to some extent interchangeable in the regressions.[23] The appropriate check is thus to compare the coefficient on local public services across specifications.

[22] This approach is also used by Nechyba and Strauss (1998).

[23] To see this argument, notice that one can always use the local budget constraint, $\tau_c \bar{y}_c = g_c$ to eliminate the tax rate from the model in equation (3). If there is some sorting on income then one should expect that the influence of the tax rate is reduced when controlling for individual income.

4.3 Is subsequent mobility greater among long-distance migrants?

The purpose of this section is to examine the patterns of subsequent mobility in the mover category. In particular we ask whether there is evidence that long-distance migrants move more within Stockholm relative to short-distance migrants in order to correct for sub-optimal initial choices.

The basic idea is the following. The two categories (short- and long-distance movers) have received an analogous prior "treatment", since both categories have moved to a new municipality in the Stockholm local labor market. We noted earlier that the response to variations in local public services was decidedly lower among long-distance movers; compare columns III and IV in Table 3. A natural explanation for this pattern may be that long-distance movers either lacked the information, or faced restrictions, such that they could not optimize with respect to local public services. More internal mobility for the long-distance movers within Stockholm after the initial move would be consistent with this hypothesis.

To provide some evidence on this issue we have extracted data from the LINDA database on the municipalities of residence between 1992 and 1997. Since we observe the place of residence only at the 31st of December each year, these data may be thought of as grouped duration data. With these data we can estimate the discrete proportional hazard associated with subsequent moves. We will only focus on the first subsequent move. We begin by sketching the relationship between the proportional hazard for continuous data and the proportional hazard for grouped data; Sueyoshi (1995) and Beck *et al.* (1998) are useful references on this topic.

Suppose that a Cox proportional hazard model can represent the hazard rate in continuous time, i.e.,

$$\lambda(t, X_{ic}, \theta) = \lambda_0(t)\exp\left(X_{ic}\theta\right) \tag{5}$$

where $\lambda_0(t)$ denotes the baseline hazard, X_{ic} a vector of independent variables, and θ a vector of parameters to be estimated. By definition the hazard rate equals $\lambda = f(\cdot)/S(\cdot) = -d\ln S(t,\cdot)/dt$, where $S(\cdot) = 1 - F(\cdot)$ denotes the survival function; therefore,

$$S(t, X_{ic}, \theta) = \Pr(T \geq t) = \exp\left[-\int_{0}^{t} \lambda(s, X_{ic}, \theta) ds\right] \tag{6}$$

where T denotes duration. Suppose that we have observed a duration until time t_{k-1} and want to ask the question: What is the probability that the duration will stop between time points t_{k-1} and t_k? If $y_{it_k} = 1$ denotes the occurrence of this event, we have: $\Pr(y_{it_k} = 1) = 1 - S(t_k, X_{ic}, \theta | T \geq t_{k-1})$. Using (5) and (6) we get

$$\begin{aligned} \Pr(y_{it_k} = 1) &= 1 - \exp\left[-\int_{t_{k-1}}^{t_k} \lambda(s, X_{ic}, \theta) ds\right] = 1 - \exp\left[-\exp(X_{ic}\theta) \int_{t_{k-1}}^{t_k} \lambda_0(s) ds\right] \\ &= 1 - \exp\left[-\exp(X_{ic}\theta + \alpha_{t_k})\right] \end{aligned} \tag{7}$$

where $\alpha_{t_k} = \ln\{\int_{t_{k-1}}^{t_k} \lambda_0(s) ds\}$.[24] Equation (7) is a model for a binary dependent variable with a "complementary log-log link". Notice that the baseline hazard is not identified from group duration data; only the change in the baseline hazard between successive time points is identified.[25]

To be more specific, we will estimate the following general model using pooled cross section data with period specific constant terms

$$\Pr(y_i = 1) = 1 - \exp\left[-\exp(X_{ic}\theta + \gamma LDM_i + t\alpha + (LDM_i \times t)\phi)\right] \tag{8}$$

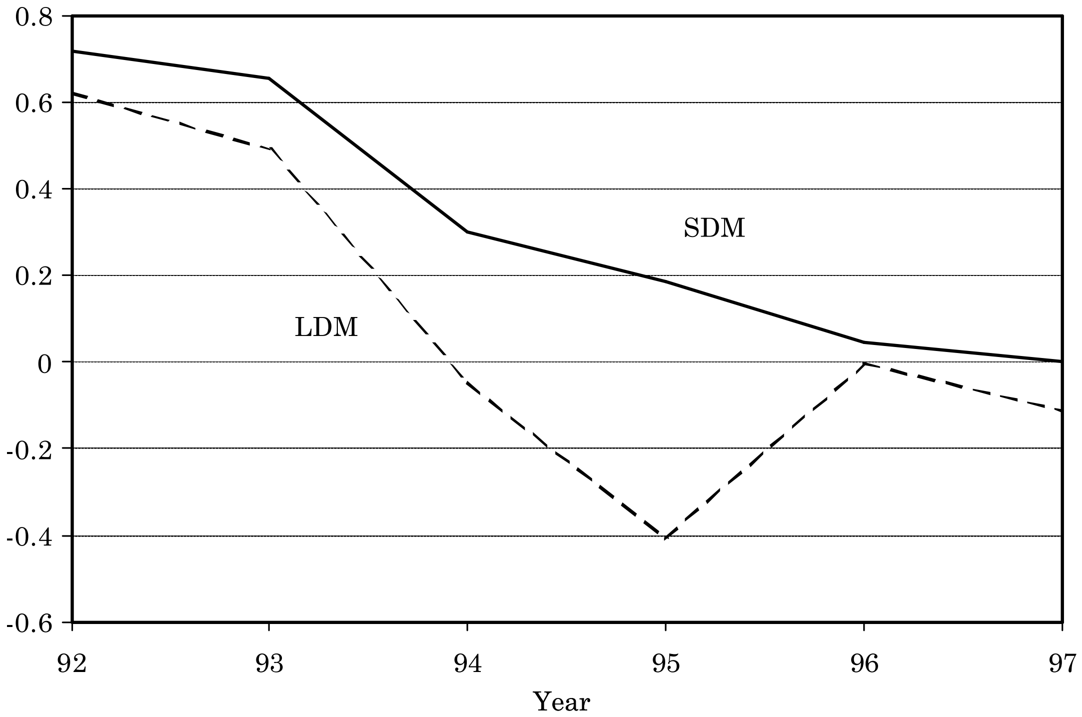
where $LDM_i = 1$ if the individual originally was a long-distance mover and t a vector of time dummies. With estimates of $\gamma, \alpha, \text{and}\, \phi$ we can test the hypothesis that long-distance movers are more likely to move again within Stockholm.

In the data there are, of course, moves to municipalities outside Stockholm. Therefore, we specify a competing risks type of model, where we treat individuals leaving Stockholm as right-censored when estimating the discrete hazard for a move within Stockholm, and vice versa.

[24] The derivation in equation (7) is valid also for time-varying X_{ic} so long as X_{ic} (as measured) does not change in the interval t_{k-1} to t_k.

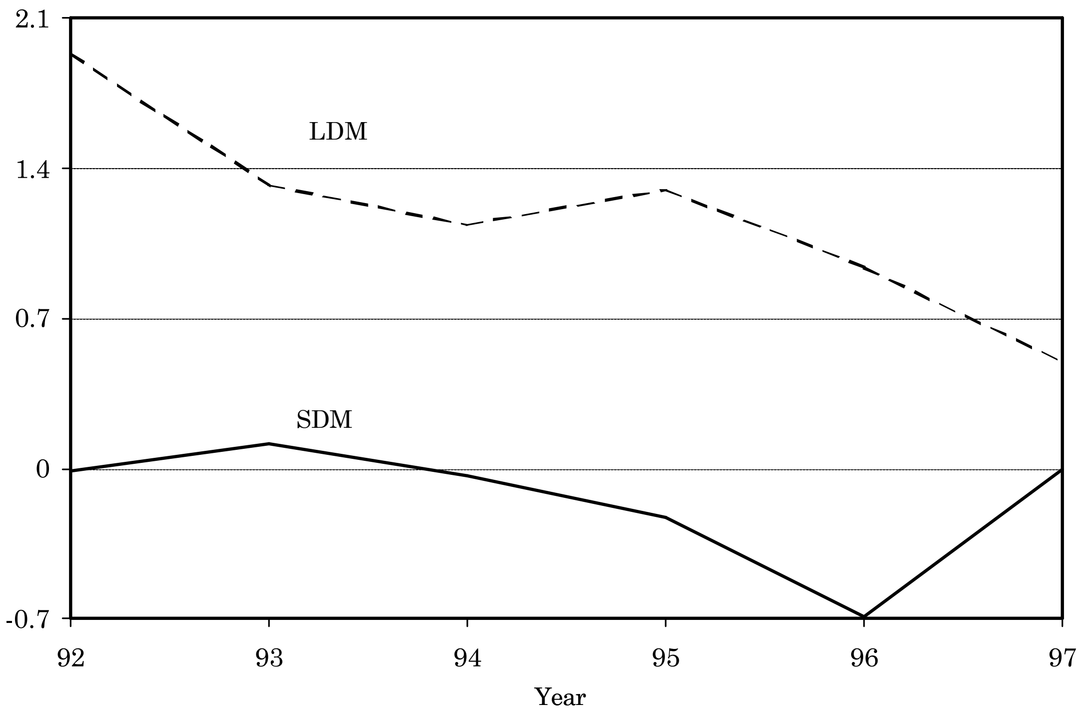
[25] Rather than postulating the proportional hazard directly, one might consider working with different assumptions about the distribution of $Z_{ic}(t) = X_{ic}\theta + \alpha(t)$. By definition, the hazard rate in a particular interval, k, equals $\lambda_k(t,\cdot) = \alpha'(t) f(Z_{ic}(t)) / [1 - F(Z_{ic}(t))]$. The proportional hazard is the result if $Z_{ic}(t)$ is Type-I extreme value distributed. The standard Logit is the result if $Z_{ic}(t)$ is logistically distributed; see Sueyoshi (1995). The logit results are very similar to those presented here.

Figure 3: Kaplan-Meier hazards for a move within Stockholm.



Note: The time effect for short-distance movers (SDM) in 1997 is normalized to 0.

Figure 4: Kaplan-Meier hazards for a move outside Stockholm.



Note: The time effect for short-distance movers (SDM) in 1997 is normalized to 0.

A simple Kaplan-Meier estimator of the hazards for the two groups conveys the basic message of the results. The Kaplan-Meier hazards for a move within Stockholm are presented in Figure 3, while the hazards for a move outside Stockholm are reported in Figure 4.

As the figures illustrate, long-distance movers are more likely to make yet another long-distance move and less likely to move within Stockholm. Thus, there seems to be no support for the hypothesis that long-distance movers re-optimize with respect to local public services.

Table 5 shows some results of the Cox regressions. We first tested whether we could impose equality of the baseline hazard for short and long-distance movers (i.e. we tested whether $\phi = 0$); this hypothesis could not be rejected. Columns I and III, therefore, present the baseline hazards when we restrict the hazards for long-distance movers to be proportional to the one for short-distance movers. Irrespective of destination, the hazards are decreasing over time. Thus the risk of subsequent mobility is greatest just after having settled in a new place of residence.

The qualitative nature of these results does not change when we add a set of control variables, as shown in columns II and IV. When individual characteristics are controlled for, the difference between the two groups is increased (reduced) when it comes to the risk of a move within (outside) Stockholm.[26] In particular, variables capturing the economic status of the individuals do a better job in the equation for moves outside Stockholm. One interpretation of this result is that the original group of long-distance movers is a selected group that is more vulnerable to economic shocks, implying that they are more likely to be subjected to a poor economic outcome and induced to make yet another long-distance move.

[26] We also considered introducing some local characteristics into the equations. The overall flavor of the results did not change when we amended the regressions with the same set of local characteristics as in column I of Table 3. Still, we formally rejected the hypothesis that the coefficients on the local characteristics were jointly zero for moves within Stockholm, although none of the individual coefficients had a z-value exceeding 1 in absolute value.

Table 5. Discrete hazard results. Competing risks specification.

	Moves within Stockholm		Moves outside Stockholm	
	I	II	III	IV
Individual characteristics				
Female		−.137		.157
		(.081)		(.137)
Age		−.057		−.090
		(.028)		(.046)
Age squared (×10^3)		.472		.962
		(.371)		(.598)
Immigrant		.117		−.501
		(.105)		(.203)
Post high school education		−.074		.003
		(.085)		(.146)
Earnings (×10^3)		-.006		−.186
		(.040)		(.081)
Unemployed		.152		.167
		(.120)		(.185)
Welfare recipient		.115		.350
		(.148)		(.213)
Household size		−.074		−.091
		(.085)		(.072)
Long-distance mover	−.193	−.296	1.449	1.337
	(.094)	(.096)	(.135)	(.143)
Time effects (1997=0)				
1992	.723	.572	.770	.622
	(.151)	(.152)	(.251)	(.255)
1993	.644	.540	.421	.313
	(.250)	(.155)	(.268)	(.269)
1994	..250	.182	.252	.175
	(.168)	(.168)	(.282)	(.283)
1995	.089	.051	.262	.213
	(.177)	(.177)	(.287)	(.288)
1996	.060	.049	−.135	−.153
	(.182)	(.182)	(.031)	(.321)
# non-censored obs	640	640	227	227
# obs	9,110	9,110	9,110	9,110

Notes: Standard errors in parentheses. All regressions include a non-reported constant. Earnings (lagged once), unemployed (lagged once), welfare recipient (lagged once), and household size are treated as time-varying characteristics.

5. Concluding remarks

In this paper we have examined whether individuals are attracted to regions offering a more attractive bundle of local public services and income tax rates. We have found a robust positive (negative) relationship between local public services (local income tax rates) and the residential choices of short-distance migrants. Moreover, we found that high-income individuals valued the aggregate basket of public services less than low-income individuals. This is in line with the observation that Swedish municipalities are responsible for important welfare programs to which high-income individuals do not qualify. When decomposing aggregate expenditures into its component parts we found that expenditures on education and “other purposes” (including, e.g., recreation and transfer programs) were the two most important items.

The evidence in favor of Tiebout related migration is much weaker – both in terms of size and significance – for those who entered from another local labor market. One plausible hypothesis tested in the paper is that long-distance movers lacked the information or the resources to make the optimal choice with respect to public services; after all, they presumably entered the Stockholm labor market mainly for labor market reasons. If so, one could expect that they should move again within Stockholm to correct for a sub-optimal initial choice. We found no evidence in favor of this hypothesis. Relative to short-distance movers, the total rate of subsequent mobility is greater among those who originally entered from another local labor market; however, this is only because they move out of Stockholm to a greater extent. To us, the evidence suggest that long-distance movers are more susceptible to economic shocks; changes in local labor market conditions forces them to make yet another long-distance move.

The fact that results differ depending on the underlying population relates to a wider methodological question. Should one base estimates of the average valuation of public services on movers or the resident population? There are pros and cons with each choice. On the one hand, one can tell plausible stories implying that (short-distance) movers are drawn from the upper end of the distribution of the valuation of public services. On the other hand, mobility costs may prevent stayers to optimize with respect to public services. It seems that the approaches taken in the previous literature is exclusively driven by data availability. As far as we know, there has been little systematic research pertaining to this question. We leave this observation as a call for future research.

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Appendix

A1. Additional estimation results

Table A1: Some additional estimates. Short-distance movers.

Variables	**I**	**II**
Expenditure (in logs)		
Total		2.678
$(\ln(g))$		(.784)
Child care	5.819	
	(1.064)	
Interaction:	–.140	
ln(Child care) ×(age)	(.030)	
Compulsory school	–3.326	
	(1.337)	
Interaction:	.109	
ln(Education) ×(age)	(.038)	
Elderly care	.264	
	(.810)	
Interaction:	–.018	
ln(Elderly care) ×(age)	(.023)	
Other purposes	.932	
	(.502)	
Interaction:	.003	
ln(Other purp.) ×(age)	(.014)	
"Prices" (in logs)		
Tax retention rate	–2.920	2.992
$(\ln(1-\tau))$	(8.786)	(6.635)
Interaction: $\ln(1-\tau)\times$ (age)	.710	
	(.240)	
Average house price	–.548	–.227
$(\ln(\rho))$	(.254)	(.338)
Marginal house price		67.342
$(\ln(p))$		(35.721)
Other variables		
Population size ($\times 10^{-5}$)	.334	.347
	(.025)	(.040)
Vacant rentals ($\times 10^{-2}$)	.523	.585
	(.154)	(.243)
Social assistance	–.005	–.002
	(.007)	(.010)
Share of foreign citizens	4.365	4.351
	(1.307)	(2.353)
Local unemployment	–88.710	–66.315
	(26.977)	(47.102)
Pseudo R-squared	0.152	0.158
Log L	–3,989	–3,963
# individuals	1,444	1,444

Notes: Standard errors in parentheses. Column I reports the results when each expenditure item is interacted with age. Column II reports results from the polytomous choice model. The latter model includes interactions between a set of community dummies and individual income. To conserve space we do not report the coefficients on these interactions; they are jointly significant, however. The construction of the marginal house price is presented in Appendix A2.

A.2 Construction of some variables

Here we describe how two of the variables used in the empirical analysis have been constructed.

Education

Teaching expenditure at the compulsory level (E_c) was missing for four municipalities. For all municipalities, however, there is information about total expenditure at the compulsory level (T_c). For the municipalities with missing observations we applied the following imputation procedure. First we calculated the share of total expenditure devoted to teaching at the compulsory level: $\gamma_c = (E_c/T_c)$. Then we averaged g_c for all municipalities where this ratio was observed, i.e., $\gamma = (\sum_c \gamma_c/22)$. Finally, we imputed teaching expenditures for the communities with missing information as: $\hat{E}_c = \gamma \times T_c$.

Marginal house price (used in the polytomous choice model)

The marginal house price for a community is the β-coefficient from the community-specific regression

$$\ln p_{ic} = \alpha_c + \beta_c SIZE_{ic} + u_{ic}$$

where $\ln p_{ic}$ is the log of the sales price of house *i* (in SEK) in community *c* in 1990 and $SIZE_{ic}$ is the size of the house (in square meters).

The data come from the "Prices of Real Estates 1990" (published by Statistics Sweden. This data source contains information about housing characteristics and sales prices for *all* real estate transactions in Sweden in 1990. The sales price is the price according to the sales contract, which must be sent to the local court in order to obtain legal confirmation of the ownership. The information about the housing characteristics comes from the form the homeowner must submit to the tax authorities.